Convergence of Japanese Local CPIs with Structural Breaks

著者名(英)	Hidehiro Ikeno
journal or publication title	駿河台経済論集
volume	21
number	2
page range	109-131
year	2012-03
URL	http://doi.org/10.15004/00000533

Articles

Convergence of Japanese Local CPIs with Structural Breaks

Hidehiro Ikeno

Abstract

This article examines whether relative local price levels converge in Japanese cities, using monthly series from 1970 to 2009. Univariate unit-root tests are used, instead of panel unit-root tests whose results involve in difficulty of economic interpretation. Without structural breaks, the null hypothesis of unit root in the relative local CPI is rejected for only a few series even by powerful unit-root tests. With structural breaks, however, the null hypothesis is rejected for many series. These results indicate that the relative local CPI is characterized in many cities as a stationary process with fast adjustment but with occasional structural changes.

Key words: unit root; structural break; CPI; convergence; purchasing power parity

JEL codes: E31; R1

1. Introduction

This article examines whether relative local price levels converge in Japanese cities, using monthly series of local consumer price indices (CPIs) from 1970 to 2009. Methodologically, univariate unit-root tests with structural breaks are used, in addition to univariate unit-root tests without structural breaks. Panel unit-root tests are not employed. The results indicate that the null hypothesis of unit root is rejected for only a few series if a structural break is not taken into account. If one or two structural breaks are admitted, however, the null hypothesis of unit root is rejected for many series. With structural breaks, the adjustment rates to the base level are rather fast, compared to results in previous studies. These results suggest that the relative local price level is characterized in many cities as a stationary process with rather fast adjustment while it faces occasional structural changes. They also suggest that failure to find the convergence in previous studies of domestic purchasing power parity (PPP) and estimation of the slow adjustment of the relative price level are likely, in part, to be due to ignoring the possibility of structural breaks.

Does the PPP hold within a single currency area? This issue is a major reason for studying the convergence of local price levels within a single country. If local markets are more integrated with each other, the PPP is more likely to hold. If local markets are not well integrated, however, the PPP does not necessarily hold. There have recently been many studies of convergence of domestic local price levels within a single country. Convergence of relative local price levels is necessary for the PPP to hold within a country. It is a typical methodology to examine whether relative local price levels follow a stationary process.

Many recent studies employ panel unit-root tests to examine stationarity, and assert rejection of the null hypothesis of unit root and hence the convergence of local price levels. Among the most representative panel unit-root tests used in the literature are the Im, Pesaran, and Shin (IPS) test (Im et al., 2003), and the Fisher test developed by Maddala and Wu (1999). Studies using panel unit-root tests emphasize that they are more powerful in rejecting the null hypothesis of unit root than univariate unit-root tests. Traditional univariate unit-root tests, such as the augmented Dickey-Fuller (ADF) test, are known for their poor power and, it is known in the literature that they scarcely reject the null hypothesis of unit root in relative local price levels.

Cecchetti et al. (2002) is a seminal study which uses panel unit-root tests in this field. Using CPIs, they conclude that local price levels in major US cities are mean-revert, but converge at a very slow rate. Esaka (2003) applies panel unit-root tests to disaggregated Japanese price data as well as the general CPI. His study concludes that the PPP holds better for tradable goods than for nontradable goods. Nagayasu and Inakura (2009) claim that the stationarity of the Japanese relative local price levels, which implies the PPP within Japan, using panel unit-root tests.

However, a panel unit-root test is not used below. A problem with the use of panel unit-root tests is the interpretation of rejecting the null hypothesis. It is inappropriate to conclude that many local prices converge by rejecting the null hypothesis in a panel containing numerous series. The null hypothesis in the IPS and Fisher tests, which are quite often used in the literature, is that all series in the panel contain a unit root. Rejection of the null hypothesis therefore implies that at least one stationary series exists. It does not imply that all or the majority of the series are stationary. Even in the case of rejection, it is possible that only a few series are stationary in the panel and that the others contain a unit root. Breuer et al. (2001) and Sonora (2009) point out the problem of this misleading inference from panel unit-root tests.[1]

This article examines the convergence of local price levels, as follows.

First, stationarity of relative local price levels is examined without assuming structural breaks. Several univariate tests are used, some of which are supposedly more powerful in rejecting the null hypothesis than most-widely used univariate unit-root tests such the ADF and Phillips-Perron tests. Those powerful tests used below are the modified Dickey-Fuller test based on generalized least-square detrending (denoted DF-GLS test) developed in Elliot et al. (1996); the tau test developed in Busetti et al. (2006); and the Horvath-Watson (HW) test developed in Horvath and Watson (1995). Sonora (2008) successfully rejects the null hypothesis of unit root in relative local price lev-

1 The Levin and Lin test (Levin and Lin, 1992) is also used in the literature as well as its modified version, the Levin, Lin, and Chu test (Levin et al., 2002). While rejection of their null hypothesis implies that all series are stationary, they assume a common unit-root process across series in the panel. Such an assumption is highly undesirable in the study of the convergence of domestic local price levels.

els by some of these unit-root tests in his study of the US economy.

Second, stationarity is examined by unit-root tests with structural breaks. It is possible that long-run relative local price levels change over time. This would cause structural breaks in the series. It has been already pointed out that unit root tests tend to fail to reject the null hypothesis if structural breaks are not taken account of; see, for example, Perron (1989). This article uses unit root tests which admit changes in the mean with endogenously estimated breakpoints, developed by Perron and Vogelsang (1992) and Clemete et al. (1998).

The results indicate that even the powerful tests fail to reject the null hypothesis of unit root in many relative local price levels if structural breaks are not taken into account. However, once structural breaks are admitted, the null hypothesis of unit root is rejected for many series. The adjustment rates of the local price levels are estimated to be rather faster than those estimated in previous studies if structural breaks are admitted. The medians of the half-lives of shocks are estimated to exceed slightly one year at most. They are often much shorter than one year. Estimated structural breaks are not distributed uniformly over the sample period. Many estimated structural breaks fall within certain periods of time, such as the period of the hyperinflation in the 1970s and the start of the "Bubble Economy" in the 1980s. These findings imply that relative local price levels converge rather fast while they are subject to occasional structural breaks, and that the structural breaks take place simultaneously in many cities.

The rest of this article is organized as follows. The second section explains the data. The third section presents preliminaries of analysis. The fourth section discusses the analysis which takes no account of structural breaks. The fifth section discusses the analysis which takes account of structural breaks. The sixth section discusses adjustment rates of the relative local price levels and estimated breakpoints in the series. The seventh section concludes.

2. Data

This article uses monthly series of the general CPI in 47 Japanese prefecture capitals from January 1970 to December 2009, except for Naha in Okinawa for which the data begin in January in 1975.[2] The national general CPI is also used over the period from January 1970 to December 2009. The period from 1970 to 2009 covers various phases of the Japanese economy. The price movement over this period ranges from high inflation to mild deflation. The early part of the sample period is characterized by high inflation due to the first and second oil shocks as well as the excessive money supply growth in the early 1970s. The later part is characterized by mild but persistent deflation, which began at the end of the 1990s. The series used take 2005 as a base. They were taken from the homepage of the Japanese Statistical Bureau, and were then seasonally adjusted by the Census X11 method for the present study.[3]

Table 1 provides overview of movements in the national and local CPIs. It shows the national inflation rate, the highest and lowest local inflation rates of the prefecture capitals and their difference over several selected periods. The entire period from 1970 to 2009 is partitioned into 20-year periods, and also into 10-year periods. Inflation rates in the table are average annualized rates expressed as a percentage. Following points are observed. First, as we

---

2 The Japanese prefecture capitals are: Sapporo, Aomori, Morioka, Sendai, Akita, Yamagata, Fukushima, Mito, Utsunomiya, Maebashi, Saitama, Chiba, Tokyo, Yokohama, Niigata, Toyama, Kanazawa, Fukui, Kofu, Nagano, Gifu, Shizuoka, Nagoya, Tsu, Otsu, Kyoto, Osaka, Kobe, Nara, Wakayama, Tottori, Matsue, Okayama, Hiroshima, Yamaguchi, Tokushima, Takamatsu, Matsuyama, Kochi, Fukuoka, Saga, Nagasaki, Kumamoto, Oita, Miyazaki, Kagoshima, and Naha. These are in the order roughly running from north to south.

3 The homepage of the Japanese Statistical Bureau is http://www.stat.go.jp. I used the seasonal adjustment program for Census X11 in the PC program RATS 7.2.

Table 1　Selected annual inflation rates (in percent points)

Periods	National CPI	Highest	City	Lowest	City	Differential
1970：01 – 2009：12*	2.90	3.07	Aomori	2.73	Sapporo	0.34
1970：01 – 1989：12*	5.48	5.70	Kyoto	5.22	Miyazaki	0.47
1990：01 – 2009：12	0.36	0.74	Aomori	0.19	Nara	0.55
1970：01 – 1979：12*	8.80	9.16	Kobe	8.46	Fukushima	0.69
1980：01 – 1989：12	2.24	2.54	Osaka	1.84	Tottori	0.70
1990：01 – 1999：12	1.01	1.37	Fukushima	0.63	Kanazawa	0.74
2000：01 – 2009：12	−0.29	0.20	Aomori	−0.59	Fukuoka	0.79

Note: *excluding Naha.

take a longer horizon, the differences between the highest and lowest local inflation rates become smaller. The differences of the 20-year periods are smaller than those of the 10-year periods, and the difference of the entire period is smaller than those of the 20-year periods. Differences decay over time. Second, the differences between the highest and lowest local inflation rates are not greatly different even when the national inflation rate rises or lowers. When the entire period is partitioned into two, the first period (1970 to 1989) is characterized by high inflation and the second period (1990 to 2009) by moderate inflation. The differences of the first and second periods fall into the same range. When the entire period is partitioned into four, the national CPI movement shifts from high inflation to mild deflation. However, the differences between the highest and lowest local inflation rates are not very different from period to period. These observations suggest that the movement of local inflation rates is restrained within a certain band over the long term. Local inflation rates do not become more diversified over time. They suggest the possibility of mean-reversion of local price levels. If local inflation rates were more diversified over time, there would be no possibility of long-term convergence of local price levels.

Cecchetti et al. (2002) report that the difference between the highest and lowest US local inflation rates over each 10-year period, beginning in 1976 and ending in 1995, are respectively, 1.29 and 1.13 percent points. Their study covers 19 cities, much fewer than the present study. Therefore, the crude observation suggests that Japanese local inflation rates are less diversified than are the US local inflation rates.

3. Preliminaries

The present study analyzes local price levels, which are expressed as a value relative to the base price level. If they follow a stationary process, then the local price levels revert to the base price level, implying convergence in the long term.

The relative price level in city i at period t is defined in log as

$$p_{i,t} \equiv ln\,(P_{i,t}/P_{b,t})$$

where $P_{i,t}$ is the CPI in city i at period t, and $P_{b,t}$ is the base CPI at period t. If p_i does not follow a stationary process, the purchasing power in city i can divert from the base level indefinitely. If p_i follows a stationary process, then the relative local CPI moves around a certain long-term level, which implies a relative PPP.

This article reports results from analyses which choose the CPI in Tokyo as the base, following previous studies on the Japanese economy such as Esaka (2003) and Nagayasu and Inakura (2009). Therefore, the number of the investigated relative local price levels is equal to 46. While a particular city is chosen as the base of the relative price in many previous studies, Chmelarova and Nath (2010) show that the results are dependent on the choice of the base city in their study of US local CPIs. I used the following four CPIs as the base and executed the same econometric exercises as for the relative local price levels; the national CPI and the local CPIs in Tokyo, Osaka, and Nagano.[4] The results from all four bases presented the following outline: The null hypothesis of unit root in the relative price level is rejected in much more cities when structural breaks are taken into account than when structural breaks are not taken into account. With the structural breaks, adjustment rates are estimated to be rather fast. Hence, this article reports results only from analyses with the Tokyo base in order to avoid presenting lengthy results.[5]

In addition to unit-root tests, it is also sensible to examine cointegration between the local CPI and the base CPI in log, i.e., $ln(P_{i,t})$ and $ln(P_{b,t})$ for investigation of the convergence. If they are cointegrated with the vector $(1, -1)$, then the relative price level in city i, $p_{i,t}$, is again characterized as reverting to the base level over the long term. This method is also adopted below.

The possibility of structural breaks should be pursued in considering domestic price level convergence. Structural breaks are not paid much attention in previous studies.[6] It is possible that failure to reject the null hypothesis of unit root in the relative price level, $p_{i,t}$, is due to neglect of structural breaks. Previous studies in empirical time series analysis, including Perron (1989), point out that neglect of structural breaks often leads to false rejection of the null hypothesis of unit root.

Even when $p_{i,t}$ is characterized as a stationary process around a certain long-term value, the long-term value may undergo occasional changes. There are various possible explanations for such changes in the CPI. The CPI covers both tradable and nontradable goods, and arbitrage of the nontradable goods is often imperfect. Various economic factors specific to each city influ-

---

4　I chose bases from both megalopolitan areas and non-megaloplitan areas. Tokyo and Osaka were chosen, because they are centers of the two largest megalopolises in Japan. Nagano was chosen, because it is roughly centrally located from the geographical point of view and does not belong to any megalopolis.

5　The results from the other bases are provided upon request to the author for interested readers.

6　Sonora (2009) shows usefulness of models with structural breaks for analysis of the domestic price convergence.

ence the prices of many nontradable goods including rents which largely affect the CPI. These factors are often subject to change over decades due to various causes, including the Samuelson-Balassa effect.[7] One should not assume away the possibility of change in values to which local price levels converge in the long run. In other words, one should not assume away the possibility that $p_{i,t}$ is characterized as a stationary process with structural changes in the mean.

4. Analysis without structural breaks

This section analyzes stationarity of the relative local CPI without taking account of structural breaks. No structural break is assumed over the sample period.

The following unit-root and cointegration tests are executed to examine stationarity;

(1) the ADF test,

(2) the DF-GLS test, developed in Elliot et al. (1996),

(3) the tau test, developed in Busetti et al (2006), and

(4) the HW test, developed in Horvath and Watson (1995).

In addition, the following stationarity test is executed;

(5) the KPSS test, developed in Kiatowski et al. (1991).

A time trend is not considered below, because the inclusion of a time trend is not compatible with the PPP.

The ADF and DF-GLS tests are executed on the following equation;

$$\Delta p_{i,t} = \mu_i + \rho_i p_{i,t-1} + \Sigma_{k=1}^{qi} \alpha_{i,k} \Delta p_{i,t-k} + e_{i,t} \tag{1}$$

where the number of lag lengths, q_i, is determined by Campbell and Perron's (1991) top-down-t approach. Starting with the number of lag lengths $k=12$, k is reduced by one each time as long as the last lag term is insignificant, i.e., the t-value of the last lag term is less than 1.645.[8] The DF-GLS test is known

7 For the Samuelson-Balassa effect, see Bahmani-Oskooee and Nasir (2005).

8 This value asymptotically corresponds to the 10% significance level.

to be more powerful than the ADF test.

Busetti et al. (2006) propose the tau test. They propose to extract the last observation from each dataset and execute the Dickey-Fuller test without a constant term in order to enhance the power. It is known that the ADF test often fails to reject the null hypothesis when the initial value of the series is distant away from the constant term in the ADF equation, and the tau test aims to overcome this problem. Below, following Busetti et al. (2006), the average of the last twelve observations is extracted from each observation, and the ADF test is then executed without the constant term. Use of the average, instead of the last observation, aims to avoid introducing noise associated with the last observation. The number of lag lengths is again determined by the top-down-t approach, starting at 12. Busseti et al. (2006) find that the tau test is useful for rejecting the null hypothesis of unit root in their study of the Italian economy.

The HW test examines the existence of cointegration between two variables. This test is claimed to be powerful by Horvath and Watson (1995). Cointegration between $ln(P_{i,t})$ and $ln(P_{b,t})$ is examined with the cointegration vector prespecified as $(1, -1)$. The null hypothesis is nonexistence of cointegration; the Wald statistic is used for the test. The number of lag lengths is determined by the top-down-t approach, starting at 12. Sonora (2008) finds that the HW tests are useful for rejecting the null hypothesis in his study of the US economy.

The KPSS test has a null hypothesis of stationarity, in contrast to the above-mentioned tests. The tests are executed with a constant term but without a time trend, and the number of lag lengths is set to eight.[9]

For all series except Naha, the sample period begins in February 1971, allowing twelve lag lengths, and the end is in December 2009. The start for

9　Even if the numbers of lag lengths are set to four and twelve with the present data set, the results from the KPSS tests are not substantially different from those stated in Table 2.

Naha is in February 1976 in all tests, because the data for Naha begin in January 1975.

Table 2 shows results from the ADF, DF-GLS, tau, HW, and KPSS tests. Critical values for ADF and DF-GLS tests are based on Fuller (1976). Critical values for tau tests are from Busetti et al. (2006), for HW tests are from Horvath and Watson (1995), and for KPSS tests are from Kiatowski et al. (1991).

Support for stationarity by the ADF, DF-GLS, tau, and HW tests, i.e., rejection of the null hypothesis of unit root or no cointegration, is minimal at any conventional significance level. Support for stationarity by the KPSS tests, i.e., acceptance of the null hypothesis of stationarity, is almost nil at any conventional significance level. The number of series for which the null hypothesis of unit root is rejected at the 10% significance level is equal to two by the ADF tests, five by the DF-GLS tests, two by the tau tests, and five by the HW tests. The null hypothesis of stationarity is rejected by the KPSS tests at the 10% significance level with all series.

The overall results deny that relative local CPIs follow a stationary process if structural breaks are not admitted. The DF-GLS and HW tests are the most successful, but even these tests do not reject the null hypothesis of unit root with many series.

5. Analysis with structural breaks

This section analyzes stationarity of relative local price levels while taking account of structural breaks.

While there are already many unit-root tests aimed at structural breaks, following are representative unit-root tests aimed at changes in the mean; the Perron and Vogelsang (PV) test, developed by Perron and Vogelsang (1992), and the Clemente, Montanes, and Reyes (CMR) test, developed by Clemente et al. (1998). The PV test is for a unit root in a time series which undergoes a single change in the mean. The CMR test is for a unit root in a time series which undergoes two changes in the mean. The latter is a two-

Table 2　Unit-root, cointegration, and stationary tests
sample period: 1971：02 – 2009：12 §

	ADF	DF-GLS	TAU	HW	KPSS
Sapporo	−1.391	0.098	−1.872	0.005	4.330***
Aomori	−0.161	−0.191	−0.920	6.384	2.079***
Morioka	−1.317	−1.168	−0.935	5.779	1.510***
Sendai	−1.948	−1.293	−1.955	8.424*	2.070***
Akita	−1.788	−1.685*	−1.793	8.246	0.662**
Yamagata	−1.408	−0.649	−1.291	5.360	2.132***
Fukushima	−1.530	−1.112	−0.965	6.930	1.166***
Mito	−1.927	−1.654*	−1.778	4.090	3.575***
Utsunomiya	−1.694	−0.789	−1.534	5.229	3.179***
Maebashi	−2.331	−0.974	−2.110	1.890	2.771***
Saitama	−1.936	−1.095	−1.978	1.622	2.699***
Chiba	−2.408	0.089	−2.622*	0.002	3.223***
Yokohama	−2.154	−1.092	−1.464	0.565	1.086***
Niigata	−1.836	−1.508	−1.650	11.269**	0.854***
Toyama	−1.986	0.005	−2.246	2.154	3.239***
Kanazawa	−1.933	0.010	−2.017	0.781	4.178***
Fukui	−1.476	−1.112	−1.516	0.136	3.763***
Kofu	−2.762*	−2.348**	−1.282	27.370**	1.094***
Nagano	−2.499	−0.952	−2.456*	11.215**	1.833***
Gifu	−2.314	−1.266	−2.279	1.507	3.326***
Shizuoka	−1.844	−0.970	−1.212	3.521	1.581***
Nagoya	−2.330	−0.236	−1.910	0.001	2.967***
Tsu	−2.092	−0.446	−1.772	8.203	0.917***
Otsu	−1.464	−1.151	−1.231	5.264	2.261***
Kyoto	−1.001	−0.220	−1.255	2.020	2.585***
Osaka	−1.052	−1.094	−0.892	4.921	3.818***
Kobe	−1.631	−1.558	−1.365	0.674	2.699***
Nara	−2.031	−1.075	−1.825	1.000	2.523***
Wakayama	−1.169	−1.140	−1.157	0.018	3.928***
Tottori	−1.825	−0.842	−1.792	4.629	2.110***
Matsue	−2.232	−2.090**	−1.737	15.084**	0.543**
Okayama	−1.438	−0.789	−1.220	1.152	2.502***
Hiroshima	−1.954	0.193	−2.236	1.933	4.195***
Yamaguchi	−0.980	−0.508	−1.100	0.633	3.694***
Tokushima	−1.305	−0.505	−1.444	2.062	4.307***
Takamatsu	−1.715	−0.708	−1.779	0.391	3.075***
Matsuyama	−1.432	−0.202	−1.576	0.105	3.542***
Kochi	−1.170	−0.404	−1.309	0.023	3.883***
Fukuoka	−2.081	−0.917	−1.921	3.861	2.642***
Saga	−1.986	−1.854*	−1.593	2.723	2.345***
Nagasaki	−1.893	−1.032	−1.553	8.131	1.431***
Kumamoto	−2.077	−0.339	−1.987	0.887	2.643***
Oita	−1.714	−0.345	−1.774	1.191	3.922***
Miyazaki	−2.134	−0.225	−2.308	0.120	3.092***
Kagoshima	−1.105	−1.101	−1.021	2.997	0.966***
Naha	−2.743*	0.296	−1.021	2.864	3.584***

Note: § This sample period does not apply to Naha, for which the sample period is 1976：02 – 2008：12. ***, **, and * represent rejection of the null hypothesis at the 1, 5, and 10% level, respectively.

break version of the former. They do not exogenously set breakpoints. They endogenously estimate breakpoints. No time trend is considered here, so that structural breaks in a time trend term are not included.

Sonora (2009) uses PV and CMR tests with the US series of relative local CPIs, and concludes that PV tests are more successful in rejecting the null hypothesis of unit root than CMR tests in his case.

The PV and CMR tests both have a null hypothesis that the times series contains a unit root. There are two models with which both tests can be executed; the innovative outlier (IO) model and the additive outlier (AO) model.

The null hypothesis of the CMR test is represented by

$$H_0: p_t = \delta_1 TB_{1t} + \delta_2 TB_{2t} + p_{t-1} + w_t$$

where $TB_{it} = 1$ if $t = T_{bi} + 1$ and 0 otherwise for $i = 1, 2$. T_{b1} and T_{b2} are the breakpoints. Hereafter, the suffix 'i' which denotes city is omitted for simplicity. The alternative hypothesis is represented by

$$H_1: p_t = c + d_1 DU_{1t} + d_2 DU_{2t} + \upsilon_t$$

where $DU_{it} = 1$ if $t > T_{bi}$ and 0 otherwise for $i = 1, 2$.

The IO model assumes that the change in the mean affects the variable gradually. There is assumed to be a transition period. The model is tested by regressing the following equation:

$$p_t = \mu + \delta_1 TB_{1t} + \delta_2 TB_{2t} + d_1 DU_{1t} + d_2 DU_{2t} + \alpha p_{t-1} + \Sigma_{j=1}^{k} c_j \Delta p_{t-j} + e_t \tag{2}$$

The test statistic t_α is the minimum value of the t-statistic which tests whether α is equal to one for all possible breakpoints.

The AO model assumes that the change in the mean takes effect instantaneously. It is examined below whether the portion removed from the deterministic part follows a unit-root process. The deterministic part is removed by regressing the following equation:

$$p_t = \mu + d_1 DU_{1t} + d_2 DU_{2t} + \tilde{p}_t.$$

The t-statistic for $\alpha = 1$, t_α, is then obtained by regressing

$$\tilde{p}_t = \Sigma_{j=0}^{k} \omega_{1j} TB_{1t-j} + \Sigma_{j=0}^{k} \omega_{2j} TB_{2t-j} + \alpha \tilde{p}_{t-1} + \Sigma_{j=1}^{k} c_j \Delta \tilde{p}_{t-j} + e_t. \tag{3}$$

The test statistic t_α is the minimum value of the t-statistic which tests whether α is equal to one for all possible breakpoints, again.

The PV test is a special case with $\delta_2 = d_2 = \omega_{2j} = 0$. The test statistic is defined as the same as in the CMR test. The IO and AO models of the PV tests are hereafter denoted by PV-IO and PV-AO. Counterparts of the CMR tests are denoted by CMR-IO and CMR-AO.

Below, the optimal lag length is based on the top-down-t approach with the referred t-value equal to 1.645, starting with the lag length set to 12. The sample period begins in March 1971 and ends in December 2009, allowing 12 lag lengths, except for Naha which starts in March 1976.

Table 3 shows results from the PV tests. This shows the test statistic t_α and the half-life of a shock and the estimated breakpoint. The half-life is a measure of persistence, and is discussed in the next section with the estimated breakpoints. Referred critical values are from Perron and Vogelsang (1992), and asymptotic critical values are referred to, considering the numbers of observations contained in the present sample periods. The number of rejections of the null hypothesis is equal to seven in the case of PV-IO and eleven in the case of PV-AO at the 10% significance level. The PV tests are more successful than any unit-root or cointegration tests which assume no structural breaks in rejecting the null hypothesis.

Tables 4 and 5 show the results of CMR-IO and CMR-AO models. They show the test statistic t_α, the half-life of a shock, and the estimated first and second breakpoints, denoted as 'breakpoint 1' and 'breakpoint 2'. Referred critical values are from Clemente et al. (1998), and asymptotic critical values are referred to, again. The number of rejections of the null hypothesis is thirteen in the case of CMR-IO, and is fifteen in the case of CMR-AO at the 10% significance level. The CMR tests are more successful in rejecting the null hypothesis of unit root than the PV tests.

Integration of the results from the PV and CMR models provides a different view from that inferred without structural breaks. The null hypothesis of unit root is now rejected at the 10% significance level with many series. With the following 17 series, the null hypothesis is rejected by at least one of

Table 3 PV tests

sample period 1971：03 – 2009：12 §

	PV-IO model			PV-AO model		
	t_α	half-life	breakpoint	t_α	half-life	breakpoint
Sapporo	−4.053	15.7	1983：04	−4.398*	7.1	1986：09
Aomori	−3.129	22.3	1998：05	−3.174	14.0	2001：04
Morioka	−1.847	33.4	1977：07	−2.079	28.5	2004：10
Sendai	−3.929	8.8	1984：08	−3.779	9.3	1984：12
Akita	−2.442	23.3	1999：04	−2.207	29.3	1992：11
Yamagata	−2.429	23.9	1981：09	−2.136	26.9	1980：07
Fukushima	−2.790	28.0	1997：01	−2.576	25.0	2002：02
Mito	−4.420*	7.6	1987：08	−4.533**	6.4	1989：06
Utsunomiya	−3.491	9.7	1994：10	−3.589	9.3	1992：12
Maebashi	−4.498**	8.8	1984：09	−5.658***	10.5	1973：03
Saitama	−4.739**	4.0	1980：01	−4.915**	3.8	1979：03
Chiba	−3.815	17.6	1974：03	−4.607**	8.3	1984：01
Yokohama	−3.740	11.0	1974：06	−3.142	13.3	1973：06
Niigata	−2.423	18.5	1984：08	−2.174	21.2	1983：08
Toyama	−3.554	9.7	1985：05	−3.233	9.0	1984：10
Kanazawa	−2.794	19.0	1983：11	−2.815	14.6	1984：04
Fukui	−3.103	11.7	1985：12	−3.324	11.2	1984：11
Kofu	−5.236***	4.3	2006：10	−5.313***	4.1	2005：10
Nagano	−3.557	9.1	1983：06	−4.025	11.1	1973：10
Gifu	−4.586**	5.9	1985：11	−4.946**	5.0	1986：06
Shizuoka	−3.238	16.7	1975：08	−3.132	19.7	1973：04
Nagoya	−3.790	13.5	1976：09	−3.396	10.5	1980：03
Tsu	−2.962	18.2	2001：09	−3.056	16.2	2003：01
Otsu	−2.473	19.2	1981：06	−2.583	18.2	1980：04
Kyoto	−2.884	16.2	1996：06	−3.158	13.4	1999：09
Osaka	−4.132	9.6	1993：05	−5.106***	5.6	1994：08
Kobe	−3.540	14.6	1980：12	−3.738	11.2	1984：12
Nara	−4.475**	9.9	1976：07	−4.525**	7.9	1979：12
Wakayama	−3.802	9.6	1986：02	−4.233*	6.4	1986：04
Tottori	−2.804	18.4	1984：10	−2.406	27.5	1979：09
Matsue	−2.738	16.0	1972：11	−2.825	15.5	2009：02
Okayama	−2.821	22.4	1981：03	−2.684	20.5	1980：07
Hiroshima	−3.353	22.4	1975：08	−3.149	14.4	1986：12
Yamaguchi	−3.841	15.0	1983：08	−3.689	9.4	1986：07
Tokushima	−3.396	20.0	1981：03	−3.140	14.5	1984：03
Takamatsu	−3.284	22.0	1974：08	−3.352	24.5	1973：01
Matsuyama	−2.931	31.2	1977：06	−2.724	18.3	1984：06
Kochi	−3.870	11.1	1982：06	−3.282	12.8	1981：06
Fukuoka	−4.305*	6.2	1987：01	−3.687	6.9	1988：03
Saga	−3.633	9.0	1987：06	−3.906	7.6	1989：11
Nagasaki	−3.758	8.8	1999：10	−4.263*	7.2	2000：10
Kumamoto	−3.449	10.1	1980：12	−3.357	9.5	1980：06
Oita	−3.465	20.3	1980：10	−3.308	11.6	1983：10
Miyazaki	−3.181	26.8	1971：09	−3.568	11.9	1984：03
Kagoshima	−2.344	20.8	1999：05	−2.165	21.6	2004：01
Naha	−3.971	13.4	1988：03	−4.007	8.6	1987：04

Note: § This sample period does not apply to Naha, for which the sample period is 1976：02 – 2008：12. ***, **, and * represent rejection of the null hypothesis at the 1, 5, and 10% level, respectively.

Table 4　CMR tests: CMR-IO model
sample period 1971：03 − 2009：12 §

	t_α	half-life	breakpoint 1	breakpoint 2
Sapporo	−4.741	6.7	1983：04	1987：12
Aomori	−4.004	14.9	1986：10	1998：05
Morioka	−3.942	9.3	1984：11	2002：10
Sendai	−5.193	5.6	1984：08	2000：03
Akita	−3.827	10.9	1986：10	1999：08
Yamagata	−4.339	8.5	1985：05	1999：10
Fukushima	−4.411	11.3	1985：07	1996：11
Mito	−5.264*	6.1	1972：11	1987：08
Utsunomiya	−4.604	6.6	1976：09	1990：05
Maebashi	−5.612**	5.5	1977：01	1986：02
Saitama	−5.810**	3.5	1977：05	1981：09
Chiba	−5.876**	5.1	1976：07	1984：08
Yokohama	−4.288	9.2	1974：06	2006：09
Niigata	−5.279*	5.0	1986：05	1997：04
Toyama	−4.688	6.3	1985：12	1998：10
Kanazawa	−3.333	12.7	1975：12	1985：02
Fukui	−3.863	8.7	1976：07	1985：12
Kofu	−5.884**	3.3	2002：07	2007：11
Nagano	−5.269*	4.7	1985：09	2004：07
Gifu	−5.382*	4.4	1982：04	1988：10
Shizuoka	−4.171	9.2	1977：08	1996：06
Nagoya	−4.143	12.1	1976：09	2006：07
Tsu	−4.598	10.6	1984：08	1994：08
Otsu	−3.535	16.6	1973：04	1976：07
Kyoto	−4.163	9.2	1977：11	1996：06
Osaka	−5.280*	6.4	1971：07	1994：10
Kobe	−4.295	11.0	1973：02	1980：12
Nara	−5.505**	7.6	1973：01	1976：07
Wakayama	−4.409	7.9	1971：08	1986：02
Tottori	−4.237	7.6	1987：11	1998：03
Matsue	−4.035	8.8	1985：07	1999：12
Okayama	−4.129	13.6	1981：03	2000：02
Hiroshima	−4.285	9.3	1975：08	1987：02
Yamaguchi	−5.103	10.3	1973：05	1983：08
Tokushima	−4.273	9.9	1981：03	1988：07
Takamatsu	−4.227	17.2	1973：03	1974：08
Matsuyama	−3.999	28.4	1973：03	1974：11
Kochi	−3.966	10.9	1982：06	2009：09
Fukuoka	−5.178	4.9	1976：03	1987：01
Saga	−5.733**	3.7	1973：12	1988：06
Nagasaki	−5.070	5.4	1988：07	2000：02
Kumamoto	−4.601	6.5	1981：02	1999：06
Oita	−4.562	9.1	1976：10	1984：08
Miyazaki	−4.024	12.0	1971：09	1985：05
Kagoshima	−4.585	6.2	1988：07	1999：05
Naha	−4.772	10.4	1988：07	2007：05

Note: § This sample period does not apply to Naha, for which the sample period is 1976：02 − 2008：12. ***, **, and * represent rejection of the null hypothesis at the 1, 5, and 10% level, respectively.

Table 5 CMR. tests: CMR-AO model
sample period 1971：03 − 2009：12 §

	t_α	half-life	breakpoint 1	breakpoint 2
Sapporo	−5.245*	4.6	1984：07	1988：11
Aomori	−4.070	8.6	1987：08	2001：04
Morioka	−4.654	5.9	1988：04	2003：05
Sendai	−5.044	4.4	1986：01	1998：07
Akita	−3.974	7.8	1989：01	1998：05
Yamagata	−5.395*	4.7	1987：08	2000：08
Fukushima	−5.326*	5.4	1985：11	1998：09
Mito	−5.299*	4.9	1974：05	1989：05
Utsunomiya	−4.736	5.7	1977：11	1993：01
Maebashi	−6.983***	5.3	1973：02	1986：09
Saitama	−6.037***	3.7	1977：02	1984：07
Chiba	−5.786**	4.6	1977：02	1984：01
Yokohama	−4.075	9.4	1973：04	2007：11
Niigata	−5.154	3.6	1987：10	1998：09
Toyama	−4.497	5.5	1985：09	1999：11
Kanazawa	−3.700	10.9	1986：11	2009：09
Fukui	−4.043	7.5	1977：11	1993：04
Kofu	−5.954**	3.4	2001：09	2009：05
Nagano	−5.147	4.2	1986：10	2005：07
Gifu	−5.445*	4.2	1981：05	1988：08
Shizuoka	−4.195	8.1	1977：04	1997：12
Nagoya	−4.107	8.3	1982：01	2005：06
Tsu	−4.472	7.3	1987：09	1997：05
Otsu	−3.761	8.8	1985：05	2002：01
Kyoto	−4.722	6.6	1978：06	1997：10
Osaka	−5.631**	4.8	1991：05	1995：11
Kobe	−4.270	9.6	1983：11	1997：12
Nara	−4.959	6.6	1975：08	1979：05
Wakayama	−4.731	5.3	1975：09	1986：04
Tottori	−4.682	5.0	1987：08	1998：07
Matsue	−4.379	6.7	1988：08	1998：11
Okayama	−4.258	7.4	1984：03	2002：01
Hiroshima	−4.794	6.0	1977：01	1988：04
Yamaguchi	−4.612	6.5	1975：11	1986：07
Tokushima	−4.170	8.4	1980：05	1990：08
Takamatsu	−3.691	14.3	1973：03	1985：03
Matsuyama	−3.654	37.4	1972：07	1973：04
Kochi	−3.723	8.8	1973：12	1985：06
Fukuoka	−5.329*	4.6	1976：02	1987：12
Saga	−5.519**	3.7	1975：07	1987：08
Nagasaki	−5.372*	4.8	1988：04	2000：08
Kumamoto	−4.443	5.7	1982：01	1998：08
Oita	−4.191	6.9	1975：09	1985：03
Miyazaki	−4.419	7.6	1988：05	1996：04
Kagoshima	−4.352	5.6	1988：02	1999：09
Naha	−4.594	6.1	1987：04	1990：02

Note: § This sample period does not apply to Naha, for which the sample period is 1976：02 − 2008：12. ***, **, and * represent rejection of the null hypothesis at the 1, 5, and 10% level, respectively.

the four models, i.e., PV-IO, PV-AO, CMR-IO, and CMR-AO, at the 10% significance level: Sapporo, Yamagata, Fukushima, Mito, Maebashi, Saitama, Chiba, Niigata, Kofu, Nagano, Gifu, Osaka, Nara, Wakayama, Fukuoka, Saga, and Nagasaki.

These results indicate that in many cities the relative local CPI follows a stationary process with one or two structural changes.

6. Half-life and breakpoints

Many previous studies of the domestic convergence of price levels conclude that adjustment of a shock to the relative price level is quite slow. Typically, the adjustment rate to the convergence is measured by the half-life of a shock which is defined as $-\ln(2)/\ln(\hat{\rho}_i + 1)$, based on Equation (1), where $\hat{\rho}_i$ is the estimate of ρ_i. Cecchetti et al. (2002) estimate the median of such half-lives as between 7 – 10 years over most sample periods in the case of the US CPIs. Dayanandan and Ralhan (2005) report, in their study on Canadian CPIs, that the medians are shorter than for the US counterparts, but are still comparable. Carrion-i-Silvestre et al. (2004) report that the median is equal to 3.6 years for Spanish CPIs. Chaudhuri and Sheen (2004) estimate 5 – 11 quarters, or 1.25 – 2.75 years, in the Australian economy. Choi and Matsubara (2007) claim, using monthly data, that the medians of the Japanese CPIs are between one and two years.

Tables 3, 4, and 5 show the estimated half-lives in months, based on the models with structural breaks. The half-life is defined as $-\ln(2)/\ln(\hat{\alpha})$, where $\hat{\alpha}$ is estimated from Equations (2) and (3), following the definition $-\ln(2)/\ln(\hat{\rho}_i+1)$ which is based on Equation (1). The medians and averages are reported in Table 6, based on the results in Tables 3 to 5. The medians are

Table 6　Half-life of a shock

	PV-IO	PV-AO	CMR-IO	CMR-AO
median (in months)	15.4	11.4	8.73	5.95
average (in months)	15.5	13.5	8.98	7.06

much shorter than their counterparts in previous studies. Naturally, the half-lives in the models with double breaks, CMR-IO and CMR-AO, are shorter than those in the models with a single break, PV-IO and PV-AO. With the CMR models, the medians are much shorter than a year. Even with the PV models, the medians are shorter than 1.3 years.

Consider now the distribution of the estimated breakpoints. Figures 1 and 2 show the distributions of the estimated breakpoints of the CMR models, based on the results shown in Tables 4 and 5. They are presented as histograms showing the year in which the estimated breakpoints fall. They show the sum of the estimated first and second breakpoints according to the CMR-IO and CMR-AO models. The two figures indicate some similarities. The estimated breakpoints are not uniformly distributed. Figure 1 shows that the largest peak is in the mid-1970s, and the second largest peak in the mid-1980s. Figure 2 shows that the largest peak is also in the mid-1970s, and the second largest peak in the second half of the1980s. Both figures show almost no breakpoints after 1990. The mid-1970s is the period of hyperinflation, due

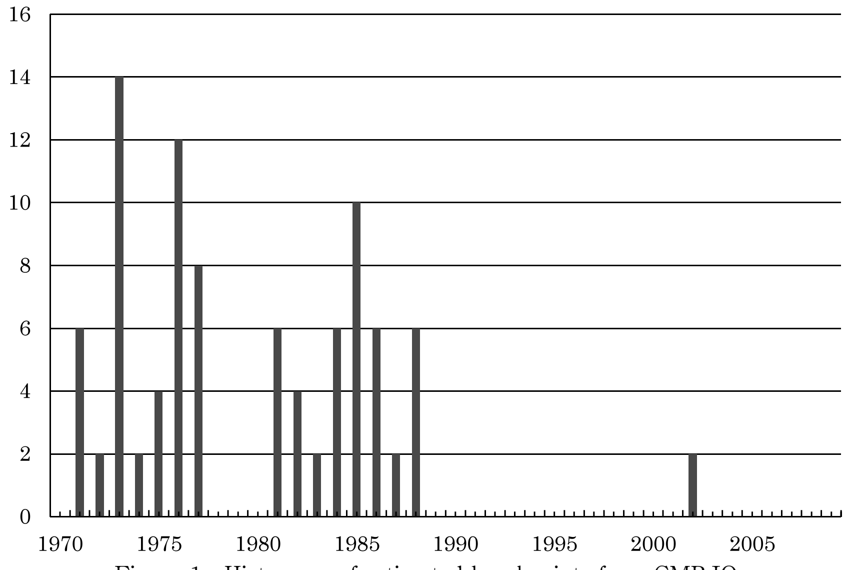

Figure 1 Histogram of estimated breakpoints from CMR-IO

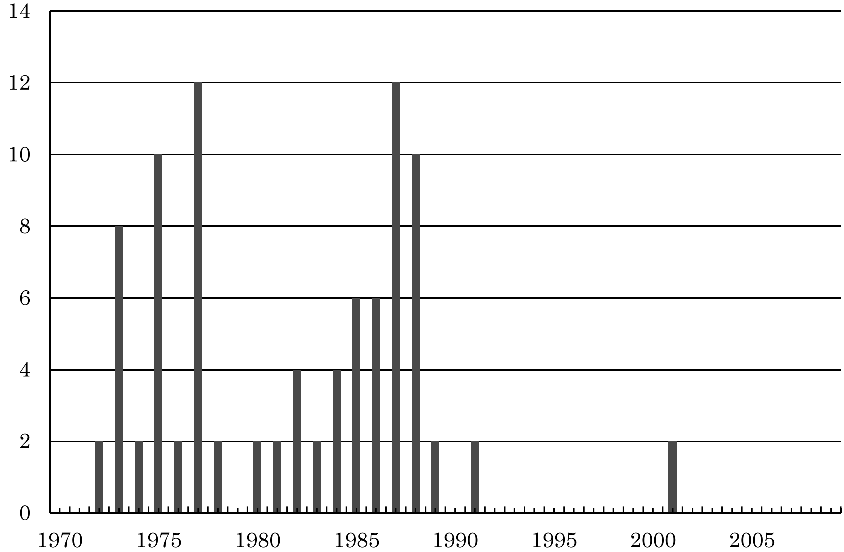

Figure 2　Histogram of estimated breakpoints from CMR-AO

to the excessive money supply growth in the early 1970s and the first oil shock. The period from the mid-1980s to the late 1980s is the start of the "Bubble Economy" in Japan.[10] These observations suggest that structural breaks in relative local price levels took place simultaneously in many cities during major national economic events.

7. Conclusions

This article shows that it is rarely successful to reject the null hypothesis of unit root in relative local CPIs without taking account of structural breaks, when analyzing the Japanese monthly data from 1970 to 2009. However, it

---

10　Choi and Matsubara（2007）assert that there was a structural break in the behavior of intercity relative price levels around 1985, based on methodology different from the one used here. They point out several possible causes of this structural break, such as changes in the distribution system within Japan which took place in this period.

also shows that the null hypothesis is rejected for many series if one takes account of structural breaks. Adjustment is rather fast according to models containing structural breaks. Structural breaks in many series are estimated to have taken place during the hyperinflation in the 1970s and at the start of the Bubble Economy in the 1980s.

It is concluded from these results that many Japanese relative local price levels are characterized as stationary process with rather fast adjustment but with occasional structural breaks. The present data set of Japanese CPIs is consistent with the claim of integration of Japanese local markets.

These results suggest that failure to reject the null hypothesis of unit root in previous studies of domestic PPP is likely to be, in part, due to failure to allow for structural breaks.

Causes of the structural breaks are not explored in this article. One of possible causes is the Samuleson-Balassa effect, which is related to the regional difference in productivity. Exploring the causes is the next promising research issue.

Acknowledgements

Earlier versions of this article were presented at the 2010 Annual Meeting of the Japan Statistical Society and at the Fall Meeting of 2010 of the Japan Society of Monetary Economics. I would like to thank Tatsuya Morisawa as well as attendants. I also would like to thank Robert Sonora, whose helpful comments and encouragement are greatly appreciated.

References

Bahmani-Oskooee, M., & Nasir, A. (2005). Productivity bias hypothesis and the purchasing power parity: a review article. *Journal of Economic Surveys, 19*, 671–696.

Breuer, J.B., McNown, R., & Wallace, M.S. (2001). Misleading inferences from panel unit-root tests with an illustration from purchasing power parity. *Review of Inter-*

national Economics, *9*, 482–493

Busetti, F., Fabiani, S., & Harvey, A. (2006). Convergence of prices and rates of inflation. *Oxford Bulletin of Economics and Statistics*, *68*, Supplement, 863–877

Campbell, J.Y., & Perron, P. (1991). Pitfalls and opportunities: what macroeconomists should know about unit roots. In O.J. Blanchard & S. Fisher (Eds.), *NBER Macroeconomics Annual 1991* (pp. 141–220). MIT Press

Carrion-i-Silvestre, J.L., Barrio, T., & Lopez-Bazo, E. (2004). Evidence on the purchasing power parity in a panel of cities. *Applied Economics*, *36*, 961–966

Cecchetti, S., Mark, N., & Sonora, R. (2002). Price index convergence among United States cities. *International Economic Review*, *43*, 1081–1099

Chaudhuri, K., & Sheen, J. (2004). Purchasing power parity across states and goods within Australia. *The Economic Record*, *80*, 314–329

Chmelarova, V., & Nath, H.K. (2010). Relative price convergence among US cities: does the choice of numeraire city matter ? . *Journal of Macroeconomics*, *32*, 405–414

Choi, C-Y, & Matsubara, K. (2007). Heterogeneity in the persistence of relative prices: what do the Japanese cities tell us ? . *Journal of the Japanese and International Economies*, *21*, 260–286

Clemente, J., Montanes, A., & Reyes, M. (1998). Testing for a unit root invariables with a double change in the mean. *Economics Letters*. *59*, 175–182

Dayanandan, A., & Ralhan, M. (2005). Price index convergence among provinces and cities across Canada: 1978–2001. *Econometrics Working Paper*, EWP0504, University of Victoria

Elliott, G., Rothenberg, T.J., & Scott, J.H. (1996). Efficient tests for an autoregressive unit root. *Econometrica*, *64*, 813–836

Esaka, T., (2003). Panel unit root tests of purchasing power parity between Japanese cities, 1960 – 1998: disaggregated price data. *Japan and the World Economy*, *15*, 233–244

Fuller, W.A. (1976) *Introduction to Statistical Time Series*, New York: John Wiley & Sons

Horvath, M., & Watson, M. (1995). Testing cointegration when some of the cointegrating vectors are prespecified. *Econometric Theory*, *11*, 984–1014

Im, K., Pesaran, M.H., & Shin, Y. (2003). Testing for unit roots in heterogeneous panels. *Journal of Econometrics*, *115*, 53–74

Kiatowski, D., Phillips, P.C.B., Schmidt, P., & Shin, Y. (1991). Testing the null hypothesis of stationarity against the alternative of a unit root. *Journal of Econometrics, 54*, 159–178

Levin, A, & Lin, C.-F. (1992). Unit root tests in panel data: asymptotic and finite-sample properties. *Working Paper*, 92–23, University of California San Diego.

Levin, A., Lin, C.-F., & Chu, C. (2002). Unit root tests in panel data: asymptotic properties and finite sample properties. *Journal of Econometrics, 112*, 81–126

Maddala, G., & Wu, S. (1999). A comparative study of unit root tests with panel data and a new simple test. *Oxford Bulletin of Economics and Statistics, 61*, 631–652

Nagayasu, J., & Inakura, N. (2009). PPP: further evidence from Japanese regional data. *International Review of Economics and Finance, 18*, 419–427

Perron, P. (1989). The great crash, the oil price and the unit root hypothesis. *Econometrica, 57*, 1361–1401

Perron, P., & Vogelsang, T. (1992). Nonstationarity and level shifts with an application of purchasing power parity. *Journal of Business and Economic Statistics, 10*, 301–320

Sonora, R.J., (2009). City relative price convergence in the USA with structural break(s). *Applied Economic Letters, 16*, 939–944

Sonora, R.J., (2008). Bivariate relative city price convergence in the United States: 1981 – 1997. *Review of Financial Economics, 17*, 92–111